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Boriss Siliverstovs*
Dierk Herzer**

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* DIW Berlin, bsiliverstovs@diw.de

** IAI Göttingen, dierkherzer@gmx.de

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DIW Berlin
Deutsches Institut für Wirtschaftsforschung
Königin-Luise-Str. 5
14195 Berlin
Tel. +49 (30) 897 89-0
Fax +49 (30) 897 89-200
www.diw.de

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Manufacturing exports, mining exports and growth: cointegration and causality analysis for Chile (1960-2001) §

Boriss Siliverstovs *

DIW Berlin

Dierk Herzer**

IAI Göttingen / DIW Berlin

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Abstract

This study examines the export-led growth hypothesis using annual time series data from Chile in a production function framework. It addresses the problem of specification bias under which previous studies have suffered and focuses on the impact of manufactured and mining exports on productivity growth. In order to investigate if and how manufactured and mining exports affect economic growth via increases in productivity, the study uses Johansen cointegration technique. The estimation results can be interpreted as evidence of productivity-enhancing effects of manufactured exports and of productivity-limiting effects of mining exports.

Keywords: Export-led growth, Chile, cointegration

JEL code: O47, F43, C32.

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*DIW Berlin, Königin-Luise Straße 5, 14195 Berlin, Germany, `e-mail: bsiliverstovs@diw.de`, tel. +49 30 89789 333, fax. +49 30 89789 108

**IAI Göttingen, Platz der Göttinger Sieben 3, 37073 Göttingen, Germany `e-mail: dierkherzer@gmx.de`, tel. +49(0)551 39 8167, fax: +49 551 39 7302

1 Introduction

One of the fundamental economic questions is how countries can achieve economic growth. One of the answers to this question relies on the export-led growth (ELG) hypothesis which postulates that export expansion, especially of manufactured goods, is a key factor in promoting economic growth. There exist a vast literature that explores the link as well as direction of causation between exports and economic growth. However, it seems that overall conclusions are, at best, mixed and contradictory (Ahmad and Kwan, 1991).

In this study, we attempt to shed an additional light on this important research topic by testing the ELG hypothesis for Chile. Chile is an interesting case study because of its recent economic history[1]. During the last four decades Chile experienced a pattern of high economic growth, which was as accompanied by a significant increase of manufactured exports both in relative and absolute terms. Chilean exports grew particularly rapidly after 1974, when a comprehensive program of economic stabilisation and restructuring was initiated. Particularly, in less than four years (1975-1979), Chile has abolished practically all quantitative import restrictions and exchange rate controls, as well as it drastically reduced imports tariffs as a part of a trade liberalisation program. Bergoeing et al. (2002) argue that these structural reforms not only significantly contributed to the export growth in the late 1970s but also these reforms have laid a sound foundation that helped the domestic economy recover from the severe economic crisis that hit most Latin American countries in 1982.

Among the few studies that have examined the causal relationship between this export performance and the Chilean economic growth, Figueroa and Letelier (1994), Amin Gutiérrez de Piñeres and Ferrantino (1997), and Agosin (1999) find evidence of export-led growth. However, these studies suffer from several methodological shortcomings: Amin Gutiérrez de Piñeres and Ferrantino (1997) can be criticised for using a simple two-variable framework in their causality test. Admittedly, causality tests are extremely sensitive to omitted variables. Even if exports are found (not) to cause growth in bivariate models, this same inference does not necessarily hold in the context of larger economic models that include other relevant variables such as capital and labour (Awokuse, 2003). Indeed, Figueroa and Letelier (1994), and Agosin (1999) estimate a larger model, but they fail to incorporate imports along with exports in their estimates. According to Riezman et al. (1996), omitting the import variable can result in spurious conclusions regarding the ELG hypothesis, because particularly capital goods imports are necessary inputs for enhancement of export and domestic production. Furthermore, export growth may relieve the foreign exchange constraint, allowing capital goods to be imported to boost economic growth. Another problem that is ignored by Figueroa and Letelier (1994), Amin Gutiérrez de Piñeres and Ferrantino (1997), and Agosin (1999) is that exports, via the national income accounting identity, are themselves a component of gross domestic product. Accordingly, exports are partly endogenous within an output equation. The

[1]For a survey on the evolution of growth and exports in Chile, see, for example, Agosin (1999).

outcome of this is a strong bias in favour of a correlation between these two variables, whatever actual causal relationship may exist between them (Greenaway and Sapsford, 1994). Finally, it should be pointed out that Figueroa and Letelier (1994), Amin Gutiérrez de Piñeres and Ferrantino (1997), and Agosin (1999) focus on 'aggregate' exports only. This may mask important differences between different export categories. Even if there is evidence in favour of the ELG hypothesis relating to certain export categories, this may not be reflected at the aggregate level, and spurious conclusions may be drawn when disaggregated exports are not examined (Ghatak et al., 1997).

The objective of this paper is to re-examine the evidence found in previous studies on the Chilean economy by carefully addressing the problematic issues pointed out above. The paper contributes to the existing literature in the following ways: First, in order to tackle the possible specification bias, we go beyond the two-variable causality relationship and estimate an export-augmented neoclassical production function. Second, we test the ELG hypothesis while controlling for capital goods imports in order to capture the role of exports in financing capital goods imports, which in turn are expected to promote growth. Third, we separate the 'economic influence' of exports on output from that incorporated into the 'growth accounting relationship' by defining the output variable net of exports. Fourth, we do not focus on total exports, but we decompose Chile's exports into its main export categories. That is to say, we examine the separate effects of mining and manufacturing exports on Chilean economic growth.

Our main finding is that we find empirical support for the ELG hypothesis in Chile with the unidirectional Granger causality running from the manufactured exports to the output but not vice versa. At the same time we record differentiated impact of the main Chilean export categories (manufactured and mining) on the output. The latter result could be interpreted as the productivity-enhancing effects of manufactured exports and as productivity-limiting effects of mining exports.

The rest of the paper is organised as follows. Section 2 discusses the theoretical background of the ELG hypothesis and derives the empirical model. In Section 3 the econometric methodology is described and the empirical results are presented. Section 4 summarises our findings.

2 Theoretical Background and Empirical Model

The export-led growth hypothesis postulates that export expansion is a key factor in promoting long-run economic growth. Several arguments can be put forward to justify the ELG hypothesis theoretically. From a demand-side perspective, it can be argued that sustained demand growth cannot be maintained in small domestic markets, since any economic impulse based on the expansion of domestic demand is bound to be exhausted quickly. Export markets, in contrast, are almost limitless and hence do not involve growth restrictions on the demand side. Thus, exports can be a catalyst for income growth, as a component of aggregate demand (Agosin, 1999).

At the same time, from a supply-side perspective, export expansion could promote economic growth through an increase in total factor productivity. First, an expansion in exports may promote specialisation in sectors in which a country has comparative advantages, and lead to a reallocation of resources from the relatively inefficient non-trade sector to the more productive export sector. Second, the growth of exports can increase productivity by offering larger economies of scale (Helpman and Krugman, 1985). Third, export growth may affect total factor productivity through dynamic spillover effects on the rest of the economy (Feder, 1983). The possible sources of these knowledge externalities include productivity enhancements resulting from increased competitiveness, more efficient management styles, better forms of organisation, labour training, and knowledge about technology and international markets (Chuang, 1998). In short, knowledge is generated through a systematic learning process initiated by exports and spilling over to the domestic economy. Fourth, export expansion may indirectly affect growth by providing the foreign exchange that allows for increasing levels of capital goods imports (Riezman et al., 1996). Increasing capital goods imports in turn stimulate output growth by raising the level of capital formation. Furthermore, recent theoretical work suggests that capital goods imports from technologically advanced countries may increase productivity and thereby growth, since knowledge and technology is embodied in equipment and machinery and therefore transferred through international trade (Chuang, 1998).

However, as it has been pointed out in the literature, it is important to distinguish between the manufactured and primary export categories when referring to their impact on the country economic performance. Thus, Lucas (1993) argues that the dynamic technological spillover effects are mainly associated with the manufactured exports rather than with the primary export. Moreover, several authors hypothesise that primary exports, and thus mining exports, could be an obstacle to greater productivity growth. The main arguments advanced in support of this hypothesis are: (i) Primary products offer no sustainable potential for knowledge spillovers, and an increase in primary exports can draw resources away from the externality-generating manufacturing sector (Sachs and Warner, 1995). (ii) Primary exports are subject to extreme price and volume fluctuations. Increasing primary exports may therefore lead to increasing GDP variability and macroeconomic uncertainty. High instability and uncertainty may, in turn, hamper efforts at economic planning and reduce the quantity as well as efficiency of investments (Dawe, 1996). As a matter of fact, Chilean economy is extremely vulnerable to fluctuating copper prices as shown in Romaguera and Contreras (1995). Against this background, we assume that the effects of exports on Chilean productivity and growth differ significantly between primary and manufactured products and therefore include both variables in our empirical analysis.

Thus, at the outset of the trade liberalisation reforms Chilean exports relied heavily on mining products, in particular copper, which accounted for about 80 percent of export income. As the results of undertaken reforms, the relative importance of the mining products has been steadily decreasing over the last decades. At the same time, the export share of manufactured goods rose from 7 percent in 1973 to about 45 percent in 2001. Although the share of mining in goods exports decreased from 90 percent in 1973, mining still accounts for 41 percent of goods exports in 2001.

Given the theoretical considerations on the possible exports role in growth promotion, we place our analysis in the framework of a simple neoclassical production function:

$$Y_t = A_t K_t^{\alpha} L_t^{\beta}, \tag{1}$$

where Y_t denotes the aggregate production of the economy at time t, and A_t, K_t, L_t are the level of total factor productivity, the capital stock, and the stock of labour, respectively. Because we want to investigate if and how manufactured and mining exports affect economic growth via increases in productivity, we assume that total factor productivity can be expressed as a function of manufactured exports, IX_t, mining exports, MX_t, capital goods imports, CM_t, and other exogenous factors, C_t:

$$A_t = f(IX_t, MX_t, CM_t, C_t) = CM_t^{\delta} IX_t^{\gamma} MX_t^{\rho} C_t, \tag{2}$$

Next we combine equation (2) with equation (1) and obtain

$$Y_t = C_t K_t^{\alpha} L_t^{\beta} CM_t^{\delta} IX_t^{\gamma} MX_t^{\rho}, \tag{3}$$

where α, β, δ, γ, and ρ are the elasticities of output with respect to K_t, L_t, CM_t, IX_t, and MX_t.

Taking natural logs, ln, of both sides of equation (3) results in the following linear function:

$$\ln Y_t = c + \alpha \ln K_t + \beta \ln L_t + \delta \ln CM_t + \gamma \ln IX_t + \rho \ln MX_t + e_t, \tag{4}$$

in which all coefficients are constant elasticities, c is a constant parameter, and e_t is the usual error term, which reflects the influence of all other factors.

It is problematic, however, that exports - via the national accounting identity - are themselves a component of output. A positive and statistically significant correlation between manufactured exports, mining exports, and aggregate output is therefore almost inevitable, even if there are no productivity effects. To remedy this problem, it is necessary to separate the 'economic influence' of exports on output from the influence incorporated into the 'growth accounting relationship'. Following Ghatak et al. (1997), we deal with this issue by using the aggregate output, net of mining and manufactured exports, NY_t ($NY_t = Y_t - IX_t - MX_t$), instead of total output, Y_t. By replacing Y_t with NY_t, we finally obtain an equation that represents the long-run relationship between the variables of interest:

$$\ln NY_t = c + \alpha \ln K_t + \beta \ln L_t + \delta \ln CM_t + \gamma \ln IX_t + \rho \ln MX_t + e_t. \tag{5}$$

This equation is estimated to determine the impact of increasing manufactured exports and mining exports on economic growth via increases in productivity. Hence, the null hypothesis that the manufactured exports does not promote growth in the long-run is given by $H_0 : \gamma = 0$. Accordingly, the ELG hypothesis is supported by the data if we find that an estimate of γ is positive and statistically significant. Similarly, the long-run effect of the primary exports on the output is given by the estimate of ρ.

3 The Econometric Approach

In our modelling we follow the general-to-specific approach advocated in Hendry and Mizon (1993) and Hendry and Juselius (2000, 2001), inter alia[2]. In particular, we start with an unrestricted VAR(p) model transformed into the error-correction form

$$\Delta x_t = \Pi x_{t-1} + \sum_{i=1}^{p-1} \Gamma_i \Delta x_{t-i} + \mu + \Phi D_t + \varepsilon_t, \varepsilon_t \sim N_n(0, \Sigma) \tag{6}$$

where μ and D_t denote a constant term and intervention dummies, respectively. Then we proceed as follows. We test for cointegration and subsequently impose the implied reduced rank restrictions on the unrestricted VAR model. Then we test for the long-run exogeneity of the system variables. We use the results of the weak exogeneity tests in order to build a parsimonious time series model for the output that passes all diagnostic tests, displays constant coefficients and possesses remarkable forecasting properties as well as in order to address the issue of causality between the output and exports variables.

The vector $x_t = (\ln NY_t, \ln K_t, \ln L_t, \ln IX_t, \ln MX_t, \ln CM_t)'$ consists of the following variables: The non-export output, NY_t, is measured by real Chilean GDP net of mining and manufactured exports. K_t is the Chilean capital stock in real terms, which was computed on the basis of accumulated capital expenditure using the perpetual inventory method. The labour variable, L_t, represents the total number of people employed each year. The variables CM_t, IX_t, and MX_t represent real imports of capital goods, real exports of manufactured goods, and real exports of mining products, respectively. All variables except L_t are measured in Chilean pesos at constant 1996 prices. The annual data span the period from 1960 till 2001. They were gathered from the *Indicadores económicos y sociales de Chile 1960-2000* and the *Boletínes mensuales* published by the Chilean Central Bank. Figure 1 shows the evolution of the variables in the period under consideration.

Since the Chilean economy has been subject to the several severe shocks during the last half of the past century, we introduce the following intervention dummies $D_t = (DI71, DI81)'$ in order to control for the large outliers in the empirical model. The former dummy variable accounts for the effects of the Allende government which persuaded inward policy, whereas the latter accounts for the effects of the recession in 1982 accompanied by the overevaluation of peso, rising international interest rates, and falling commodity prices. These intervention dummies $DIxx$ take value of 1 in $19xx$ and -1 in $19xx + 1$ and zero otherwise. As discussed in Hendry and Juselius (2000, 2001), such form of the intervention dummies ensures that the asymptotic critical values of the cointegration test remains intact.

At the first stage we determine the lag order of the VAR(p) model in equation (6) by means of the sequential modified likelihood ratio (LR) test, discussed in Lütkepohl (1991). Given rather large number of explanatory variables $n = 6$ for a given sample size $T = 42$, we allow for maximum of three lags in order to allow for sufficient degrees of freedom in our testing procedure. The results of the LR test procedure

[2] All computations and graphics has been made with PcGive 10.1 and GiveWin 2.20, see Doornik and Hendry (2001a,b).

are reported in the upper panel of Table 1. As seen, the lag length $p = 2$ is selected.

Table 1 reports the univariate as well as the multivariate misspecification tests of the chosen VAR(2) model. As seen, the VAR(2) model adequately describes the data. Even though the univariate tests report single rejections of the null hypotheses of no residual autocorrelation and normality of the residuals at the 5% significance level, we cannot reject the null hypothesis that the residuals are multivariate Gaussian white noise at the conventional significance levels. In addition, the residuals seem to have no ARCH(1) effects.

Furthermore, against the background of the several economic crises that hit the Chilean economy, it is important to check whether the empirical model displays constant parameters. The structural stability of the estimated system is checked with the three versions of the system Chow tests: 1-step, Break-point, and Forecast Chow test, displayed in Figure 2. The graphs display the recursive test statistics scaled by the respective 1% critical values. We find no evidence of structural instability of our model and proceed further by addressing its cointegration properties.

After having found the adequate unrestricted model, the next step is to proceed imposing restrictions on that model. Hence, we address the cointegration rank of the estimated system. We use the Johansen Full Information Maximum Likelihood (FIML) procedure for this purpose. Table 2 reports the results of the trace and λ-max tests. As seen, both tests indicate the presence of one cointegrating relation in the system. This conclusion is supported by the magnitude of the modulus of the largest 6 eigenvalues of the companion matrix reported for the unrestricted and restricted $r = 1$ models. As seen, there are 5 eigenvalues that are quite close to the unity in the unrestricted model and the sixth eigenvalue takes the values of 0.700 and of 0.730 in the unrestricted and restricted models, respectively.

Thus we impose the cointegration rank $r = 1$ on the system (6) and proceed with testing for stationarity, long-run exclusion, and long-run weak exogeneity of the variables. The test of stationarity of the variables in the model has been suggested in Johansen and Juselius (1992). This is a multivariate version of the Augmented Dickey-Fuller test with the null hypothesis of stationarity rather then non-stationarity. Since a linear combination of I(1) variables that is I(0), or I(0) variables themselves, could only belong to the cointegration space, it investigates whether any of the variables alone belong to the cointegration space. This test has an asymptotic χ^2 distribution with the $(p - r) = 5$ degrees of freedom.

The test for the long-run exclusion (Johansen and Juselius, 1992) investigates whether any of the variables can be excluded from a cointegrating vector. This test has an asymptotic χ^2 distribution with the $r = 1$ degrees of freedom. Finally, test for the long-run weak exogeneity investigates whether the dependent variables adjust to the equilibrium errors, represented by a cointegrating relation.

Tables 3 and 4 report the results of the tests for (trend-)stationarity and long-run exclusion, performed on the matrix of the long-run coefficients, and the tests for long-run weak exogeneity, performed on the matrix of the adjustment coefficients, respectively. According to the stationarity test, the null hypothesis that each variable is I(0) or I(0) around a linear deterministic trend is decisively rejected. Moreover, the results of the test for long-run exclusion suggest that none of the variables could be excluded from the long-run relation. This is the important result, which implies that all the variables, considered in the present study, are relevant for modelling the long-run relationship, and hence any other model based on a smaller subset of these variables would suffer from the omitted variable bias problem.

According to the univariate long-run weak exogeneity test results (see the upper panel of Table 4), we can accept the null hypothesis that the variable $\ln IX_t$, that represents the manufactured exports, is weakly exogenous at any conventional significance level. At the same time, the null hypothesis of long-run weak exogeneity of the output variable $\ln NY_t$ is decisively rejected. In addition, observe that according to the univariate test results, we cannot reject the null hypothesis that the variables $\ln L_t$ and $\ln CM_t$ are weakly exogenous only at the 1% significance level. Nevertheless, as the joint long-run weak exogeneity test results suggest (reported in the lower panel of Table 4), we cannot reject the null hypothesis that the three variables $\ln L_t, \ln IX_t, \ln CM_t$ are weakly exogenous at the 10% significance level. Imposition of further zero restrictions on the adjustment coefficients yields the following group of the four variables $\ln K_t, \ln L_t, \ln IX_t, \ln CM_t$ that are weakly exogenous either at the 5% or 1% significance levels. In order to check, whether this result is robust to the change in the sample size, we report the value of the recursive test statistics of the latter null hypothesis, scaled by the 1% critical value, in Figure 3. Observe, that the restriction that the four variables $\ln K_t, \ln L_t, \ln IX_t, \ln CM_t$ are weakly exogenous with respect to the long-run parameter values is accepted for all sample sizes. Hence, this restriction seems to be reasonable, and in our further analysis we treat these four variables as weakly exogenous with respect to the long-run parameters.

Imposing the long-run weak exogeneity restrictions on the $\ln K_t, \ln L_t, \ln IX_t, \ln CM_t$ variables results in the following cointegrating vector (displayed in Figure 4) with absolute t-values reported in parentheses below the coefficient estimates

$$\ln NY_t = \underset{(11.03)}{0.695} \ln K_t + \underset{(8.07)}{0.675} \ln L_t + \underset{(4.00)}{0.047} \ln IX_t - \underset{(5.57)}{0.333} \ln MX_t + \underset{(5.39)}{0.164} \ln CM_t \tag{7}$$

Observe that all the coefficient estimates have an expected signs and all estimates are significantly different from zero, which conform with the results of the long-run exclusion restriction tests reported in Table 3 above.

Observe that our estimation results provide empirical support for the ELG hypothesis in Chile as the coefficient estimate of γ, that measures the long-run influence of the manufactured exports on the output, is positive and statistically significant, $\widehat{\gamma} = 0.047$. At the same time, we record the negative and also statistically significant coefficient estimate of ρ, that measures the long-run impact of the mining exports on the output. This result is in line with the theoretical consideration on the negative role that primary exports could exert on the economic performance of a country as discussed above and it complies with the results reported in Ghatak et al. (1997) who tested the ELG hypothesis for Malaysia.

As expected, the capital and labour stock as well as the capital goods imports contribute positively to the output. An interesting point is that the estimate of the capital elasticity is rather high. This finding of a relatively high capital elasticity may be in line with economic theory that suggests that opening to trade and the elimination of distortions increase the average quality of capital and improve the allocation of capital towards sectors with higher marginal productivity. In addition observe, that the sum of capital and labour elasticities is greater than one, which might indicate the presence of increasing returns to scale. The existence of increasing returns to scale can theoretically explain the exceptionally high growth rates of the Chilean economy observed in the late seventies and in 1985 - 1997.

As shown in Johansen (1992), the status of long-run weak exogeneity of some variables allows us to reformulate the model (6) in terms of a conditional model, where we condition on the current and past values of the weakly exogenous variables, and the marginal models for these weakly exogenous variables. The conditional model involves two variables, $\ln NY_t$ and $\ln MX_t$, and for the rest of the variables $\ln K_t, \ln L_t, \ln IX_t, \ln CM_t$ we have marginal models, that does not include the error-correction term. Our main focus, however, is on modelling the output $\ln NY_t$ variable. Therefore we report the results of the conditional model only for this variable.

The estimated conditional model for $\ln NY_t$ with absolute t-values reported in parentheses below the coefficient estimates is

$$\begin{aligned}
\Delta \ln NY_t \quad = \quad & \underset{(6.96)}{2.264}\ \Delta \ln K_t \ - \ \underset{(4.30)}{1.406}\ \Delta \ln K_{t-1} \ - \ \underset{(4.83)}{0.090}\ \Delta \ln CM_{t-1} \\
& + \underset{(4.96)}{0.656}\ \Delta \ln L_t + \ \underset{(4.15)}{0.491}\ \Delta \ln L_{t-1} - \ \underset{(5.22)}{0.490}\ ec_{t-1} + \ \underset{(5.03)}{0.405} \\
& \widehat{\sigma} = 0.020,\ R^2 = 0.894,\ T = 40,\ F_{AR(1-2)}(2,31) = 1.94[0.160], \\
& F_{ARCH(1)}(1,31) = 0.50[0.486],\ \chi^2_{Norm}(2) = 0.91[0.634], \\
& F_{Het}(12,20) = 0.50[0.887],\ F_{RESET}(1,32) = 3.17[0.084]
\end{aligned} \tag{8}$$

The conditional model (8) is parsimonious, it has very good explanatory power, its coefficients are well determined, and the diagnostic tests show no signs of misspecification. Observe that the error-correction term is highly significant and it has the expected sign.

In addition, the recursive Chow test statistics also indicate parameter constancy (see Figure 5). This conclusion is also reenforced by the recursively calculated coefficients shown in Figure 6 along with the one-step ahead residuals. The remarkable stability of the coefficients is also supported by the ability of the model to produce accurate 1-step ahead forecasts of the dependent variable $\Delta \ln NY_t$ displayed in Figure 7 for years 1982-2001. The corresponding parameter constancy forecast tests are Forecast $\chi^2(20) = 19.54[0.487]$ and Chow $F(20, 13) = 0.79[0.684]$. In particular, notice the ability of the model to predict the recession in 1982.

Since the existence of cointegration between the variables implies existence of either unidirectional of bidirectional Granger causality, the next step is to address the direction of the causality between the output on the one hand and each of the manufactured and mining exports on the other.

In determining the Granger causality direction we can use information from the long-run weak exogeneity tests reported in Table 4 above. Recall that the long-run weak exogeneity hypothesis was rejected for both variables $\ln NY_t$ and $\ln MX_t$. This results leads to the following conclusions. First, there is a bidirectional Granger causality between these two variables. Second, the manufactured exports, $\ln IX_t$, Granger causes the output variable, $\ln NY_t$, as well.

In order to determine whether there is Granger causality from the output to the manufactured exports, we estimate the following marginal model for the latter variable with absolute t-values reported in parentheses below the coefficient estimates

$$\begin{aligned}
\Delta \ln IX_t \quad = \quad & \underset{(3.96)}{0.406}\ \Delta \ln IX_{t-1} + \underset{(0.003)}{0.003}\ \Delta \ln NY_{t-1} + \underset{(0.52)}{0.558}\ \Delta \ln K_{t-1} \\
& + \underset{(0.41)}{0.106}\ \Delta \ln MX_{t-1} - \underset{(1.68)}{0.262}\ \Delta \ln CM_{t-1} + \underset{(1.23)}{1.456}\ \Delta \ln L_{t-1} \\
& - \underset{(0.06)}{0.003} + \underset{(4.29)}{0.608}\ \mathrm{D64}_t + \underset{(6.73)}{1.092}\ \mathrm{D72}_t - \underset{(2.59)}{0.362}\ \mathrm{D74}_t - \underset{(3.20)}{0.581}\ \mathrm{D81}_t \qquad (9) \\
& \widehat{\sigma} = 0.135,\ R^2 = 0.0.781,\ T = 40,\ F_{AR(1-2)}(2,27) = 0.17[0.845], \\
& F_{ARCH(1)}(1,27) = 0.33[0.566],\ \chi^2_{Norm}(2) = 2.69[0.260], \\
& F_{Het}(16,12) = 0.80[0.662],\ F_{RESET}(1,28) = 1.01[0.322]
\end{aligned}$$

The model (9) displays no signs of misspecification and no parameter instability as seen in Figures 8 and 9. Observe that in order to fulfil the residual normality assumption we have included the impulse dummies Dxx which take value of 1 in $19xx$ and zero otherwise, which are absent in the conditional model (8). Hence, the conclusion on whether the output variable, $\ln NY_t$ Granger causes the manufactured exports can be based on the corresponding t-statistic value of 0.003 which yields the p-value of 0.997. This result suggests that the output variable does not Granger cause the manufactured exports in Chile.

4 Conclusions

This paper uses Johansen cointegration technique to examine the productivity effects of manufactured and mining exports in the context of the export-led growth hypothesis. We test the ELG hypothesis for Chile using the using the time series data for 1960 - 2001. In our analysis we employ a simple framework of an augmented neoclassical production function where total factor productivity is assumed to be a function of mining and manufactured exports, and capital goods imports.

In our study we address several aspects that have been overlooked in the previous literature that tested the ELG hypothesis on the Chilean data. First, we employ the multivariate analysis which goes beyond the two-variable causality relationship between the output and exports in estimating export-augmented neoclassical production function. Second, we test the ELG hypothesis while controlling for capital goods imports in order to capture the role of exports in financing capital goods imports, which in turn are expected to promote growth. Third, we separate the 'economic influence' of exports on output from that incorporated into the 'growth accounting relationship' by defining the output variable net of exports. Fourth, we do not focus on total exports, but we decompose Chile's exports into its main export categories, i.e. manufactured and mining exports.

Our main finding is that the manufactured exports Granger causes output but not vice versa which supports the export-led growth hypothesis for Chile. At the same time, our results indicate bidirectional Granger causality between the non-export GDP and mining exports. We also record the differentiated impact that mining and manufactured exports exert on the aggregate output. This justifies our decision to split the Chilean exports in its main categories. In connection with the theoretical foundations underpinning our model, this estimation result can be interpreted as evidence of productivity-enhancing effects of manufactured exports and of productivity-limiting effects of mining exports. The latter may be due to the problem of fluctuating commodity export prices and earnings, especially copper prices, which is well known in the Chilean literature. Romaguera and Contreras (1995), for example, found that copper price volatility had negative effects on Chilean GDP growth. Our results are consistent with the fact that manufactured exports might offer greater potential for knowledge spillovers and other positive externalities than traditional primary exports.

Accordingly, the primary conclusion that emerges from this study is that while mining and manufactured export earnings certainly contributed to the Chilean national income, exports of manufactured products have been especially important for productivity and thus for long-run economic growth. This conclusion has crucial policy implications. It is particularly important to promote exports of manufacturing goods - by avoiding trade-distorting measures that would counteract the comparative advantages, and building new comparative advantages and export opportunities in the Chilean manufacturing sector.

Finally, our findings suggest that there exists a long-run relationship between capital, labour, capital goods imports, manufactured exports, mining exports on the one hand and non-export GDP on the other. We use this cointegrating relationship in order to build an error-correction model for the Chilean net-of-exports GDP that displays constant parameters over the estimated sample and remarkable ex-post forecasting ability.

References

Agosin, M. (1999). Trade and growth in Chile. *Cepal Review 68*, 79–100.

Ahmad, J. and A. C. C. Kwan (1991). Causality between exports and economic growth. *Economic*

Letters 37, 243–248.

Amin Gutiérrez de Piñeres, S. and M. Ferrantino (1997). Export diversification and structural dynamics in the growth process: The case of Chile. *Journal of Development Economics 52*(2), 375–391.

Awokuse, T. O. (2003). Is the export-led growth hypothesis valid for Canada? *Canadian Journal of Economics 36*(1), 126–136.

Bergoeing, R., P. J. Kehoe, T. J. Kehoe, and R. Soto (2002). Policy-driven productivity in Chile and Mexico in the 1980s and 1990s. NBER working Paper 8892.

Chuang, Y.-C. (1998). Learning by doing, the technology gap, and growth. *International Economic Review 39*(3), 697–721.

Dawe, D. (1996). A new look at the effects of export instability on investment and growth. *World Development 24*(12), 1905–1914.

Doornik, J. A. and D. F. Hendry (2001a). *GiveWin: An Interface to Empirical Modelling.* London: Timberlake Consultants Press.

Doornik, J. A. and D. F. Hendry (2001b). *Modelling Dynamic Systems Using PcGive*, Volume II. London: Timberlake Consultants Press.

Feder, G. (1983). On exports and economic growth. *Journal of Development Economics 12*(1), 59–73.

Figueroa, L. and L. Letelier (1994). Exportaciones, orientación al comercio y crecimiento: Un enfoque de cointegración. *Cuadernos de Economia 31*(94), 401–421.

Ghatak, S., C. Milner, and U. Uktulu (1997). Exports, export composition and growth: Cointegration and causality evidence for Malaysia. *Applied Economics 29*, 213–223.

Greenaway, D. and D. Sapsford (1994). What does liberalisation do for exports and growth? *Weltwirtschaftliches Archiv 130*(1), 152–174.

Helpman, E. and P. Krugman (1985). *Market Structure and Foreign Trade.* Cambridge: MIT Press.

Hendry, D. F. and K. Juselius (2000). Explaining cointegration analysis. Part 1. *The Energy Journal 21*(1), 1–42.

Hendry, D. F. and K. Juselius (2001). Explaining cointegration analysis. Part 2. *The Energy Journal 22*(1), 75–120.

Hendry, D. F. and G. E. Mizon (1993). Evaluating econometric models by encompassing the VAR. In P. C. Phillips (Ed.), *Models, Methods and Applications of Econometrics*, Basil. Blackwell.

Johansen, S. (1992). Testing weak exogeneity and the order of cointegration in UK money demand. *Journal of Policy Modeling 14*, 313–334.

Johansen, S. and K. Juselius (1992). Testing structural hypotheses in a multivariate cointegration analysis of the PPP and the UIP for the UK. *Journal of Econometrics 53*, 211–244.

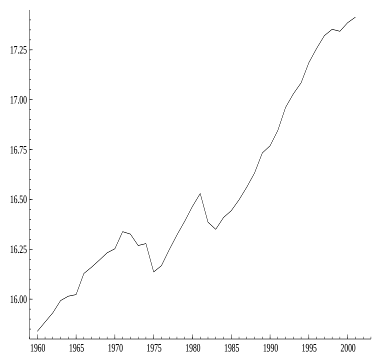
Lucas, R. E. (1993). Making a miracle. *Econometrica 61*(2), 251–272.

Lütkepohl, H. (1991). *Introduction to Multiple Time Series Analysis*. Berlin: Springer-Verlag.

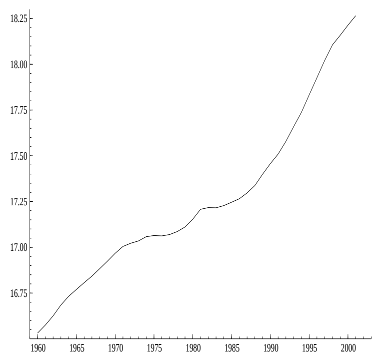
Riezman, R. G., P. M. Summers, and C. H. Whiteman (1996). The engine of growth or its handmaiden? A time series assessment of export-led growth. *Empirical Economics 21*(1), 77–113.

Romaguera, P. and D. Contreras (1995). Impacto macroeconomico de la inestabilidad del precio del cobre en la economía Chilena. In *La Administración de Los Ingresos Por Exportaciones Mineras En Bolivia, Chile Y Perú*, pp. 171–196. GRADE/Lima, Perú.

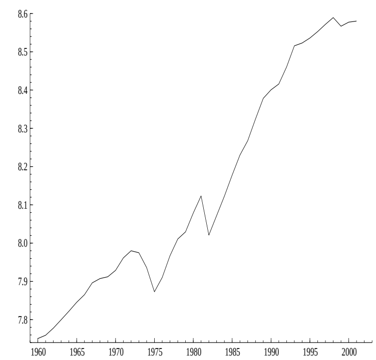
Sachs, J. and A. Warner (1995). Natural resource abundance and economic growth. National Bureau of Economic Research Working Paper 5398, Cambridge, Massachusetts.



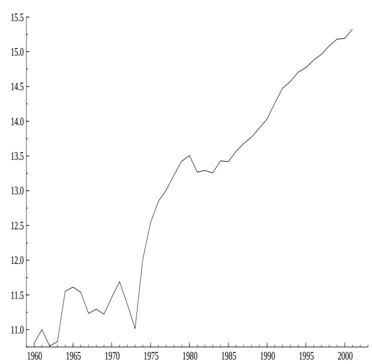
(a) $\ln NY_t$



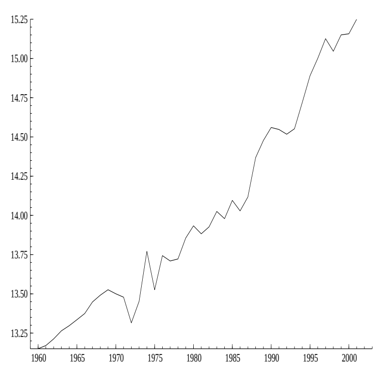
(b) $\ln K_t$



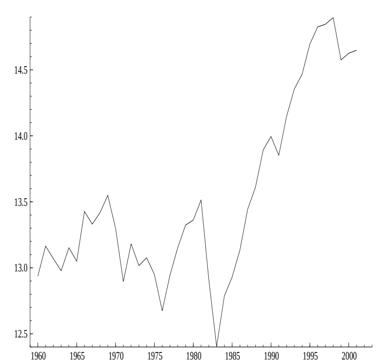
(c) $\ln L_t$



(d) $\ln IX_t$



(e) $\ln MX_t$



(f) $\ln CM_t$

Figure 1: Data: 1960-2001

Table 1: Specification tests

VAR lag selection: modified LR sequential procedure						
Lag length, p	0.00	1.00	2.00	3.00		
Loglik	110.41	408.90	466.25	505.96		
$\chi^2(36)$	na	459.23[0.000]**	70.58[0.001]**	36.65[0.438]		
Multivariate tests						
$F_{AR(1-2)}(72,49)$	=	1.33	[0.139]			
$\chi^2_{Norm}(12)$	=	16.96	[0.151]			
Univariate tests						
	$\ln NY_t$	$\ln K_t$	$\ln L_t$	$\ln IX_t$	$\ln MX_t$	$\ln CM_t$
$F_{AR(1-2)}(2, 23)$	0.78	0.59	2.63	0.40	4.73	0.13
	[0.470]	[0.562]	[0.093]	[0.676]	[0.019]*	[0.876]
$\chi^2_{DH}(2)$	1.98	2.10	6.73	5.13	0.77	0.04
	[0.372]	[0.348]	[0.034]*	[0.077]	[0.681]	[0.980]
$F_{ARCH(1)}(1, 23)$	0.72	0.052	3.61	0.15	0.99	0.24
	[0.404]	[0.820]	[0.070]	[0.700]	[0.328]	[0.630]
$\hat{\sigma}$	0.035	0.010	0.024	0.240	0.067	0.158

Table 2: Cointegration test

					Modulus of 6 largest roots	
rank	Trace test	[Prob]	Max test	[Prob]	Unrestricted	r=1
0.00	111.02	[0.002]**	44.51	[0.011]*	0.983	1.000
1.00	66.52	[0.087]	24.50	[0.433]	0.922	1.000
2.00	42.02	[0.159]	22.49	[0.202]	0.922	1.000
3.00	19.52	[0.467]	11.09	[0.647]	0.864	1.000
4.00	8.43	[0.428]	8.01	[0.386]	0.864	1.000
5.00	0.41	[0.519]	0.41	[0.519]	0.700	0.730

Table 3: Tests for stationarity, long-run exclusion

$\ln NY_t$	$\ln L_t$	$\ln K_t$	$\ln MX_t$	$\ln IX_t$	$\ln CM_t$	trend	$\chi^2(v)$	
Stationarity								
·	0	0	0	0	0		41.30	[0.000]**
0	·	0	0	0	0		39.14	[0.000]**
0	0	·	0	0	0		41.65	[0.000]**
0	0	0	·	0	0		40.69	[0.000]**
0	0	0	0	·	0		38.95	[0.000]**
0	0	0	0	0	·		41.44	[0.000]**
Trend-stationarity								
·	0	0	0	0	0	·	28.47	[0.000]**
0	·	0	0	0	0	·	33.19	[0.000]**
0	0	·	0	0	0	·	32.26	[0.000]**
0	0	0	·	0	0	·	31.22	[0.000]**
0	0	0	0	·	0	·	29.27	[0.000]**
0	0	0	0	0	·	·	29.51	[0.000]**
Long-run exclusion								
0	·	·	·	·	·		13.06	[0.000]**
·	0	·	·	·	·		12.31	[0.000]**
·	·	0	·	·	·		70.57	[0.008]**
·	·	·	0	·	·		67.24	[0.009]**
·	·	·	·	0	·		90.49	[0.002]**
·	·	·	·	·	0		12.76	[0.000]**

Notes: '0' denotes the zero restriction on the coefficient of the corresponding variable, '·' denotes unrestricted coefficient in the 6×1 cointegration vector when testing for the stationarity and long-run exclusion and 7×1 cointegration vector when testing for trend-stationarity of the variables.

The number of degrees of freedom v in the χ^2 tests corresponds to the number of zero restrictions imposed.

Table 4: Tests for long-run weak exogeneity

$\ln NY_t$	$\ln L_t$	$\ln K_t$	$\ln MX_t$	$\ln IX_t$	$\ln CM_t$	$\chi^2(v)$	
Long-run weak exogeneity							
0	·	·	·	·	·	16.30	[0.000]**
·	0	·	·	·	·	4.43	[0.035]*
·	·	0	·	·	·	7.71	[0.005]**
·	·	·	0	·	·	9.75	[0.001]**
·	·	·	·	0	·	0.01	[0.965]
·	·	·	·	·	0	4.05	[0.044]*
·	0	·	·	0	0	6.25	[0.100]
0	0	·	·	0	0	19.54	[0.001]**
·	0	0	·	0	0	8.240	[0.083]
·	0	·	0	0	0	13.77	[0.008]**

Notes: '0' denotes the zero restriction on the coefficient of the corresponding variable, '·' denotes unrestricted coefficient in the 6×1 vector of the adjustment coefficients.

The number of degrees of freedom v in the χ^2 tests corresponds to the number of zero restrictions imposed.

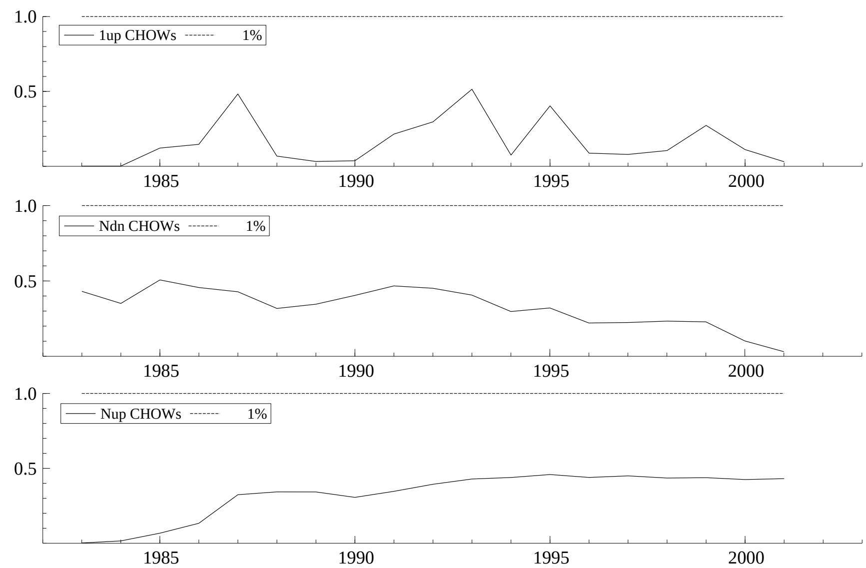


Figure 2: System Chow test statistics scaled by the corresponding 1% critical values

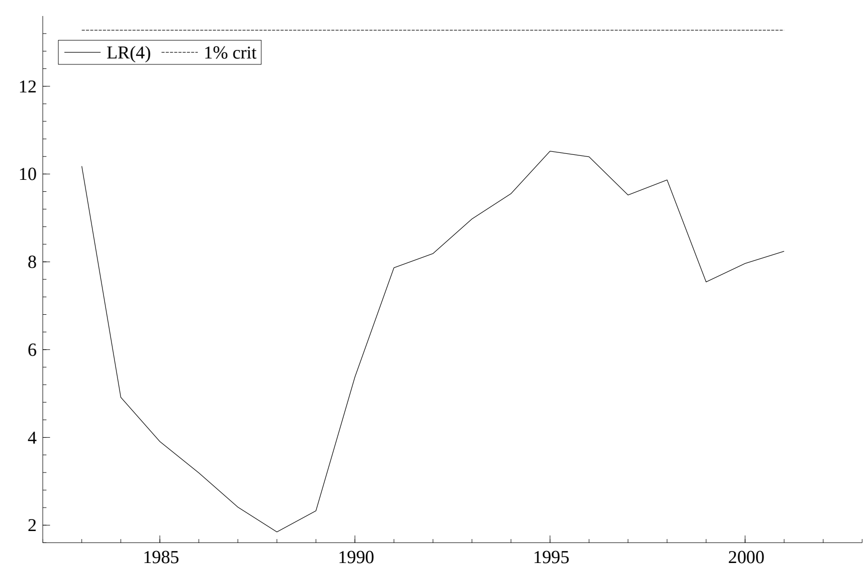


Figure 3: Recursive test statistic for long-run weak exogeneity of $\ln K_t, \ln L_t, \ln IX_t, \ln CM_t$ scaled by the 1% critical values

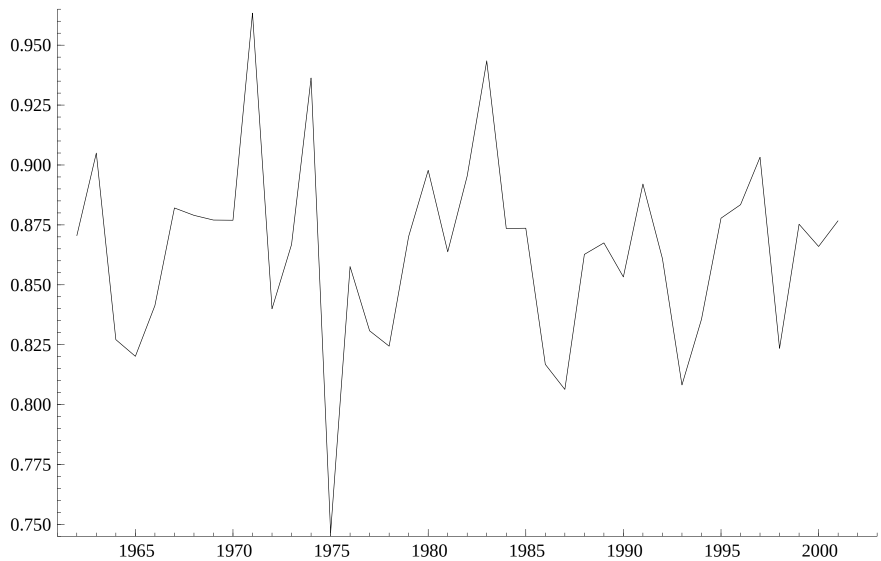


Figure 4: Cointegrating relation, equation (7)

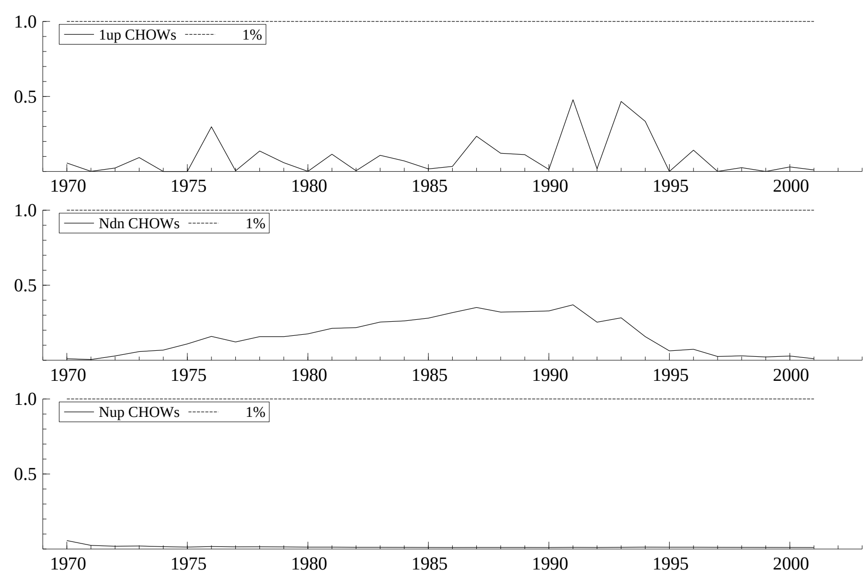


Figure 5: The recursive Chow test statistics for the conditional model (8) scaled by the corresponding 1% critical values

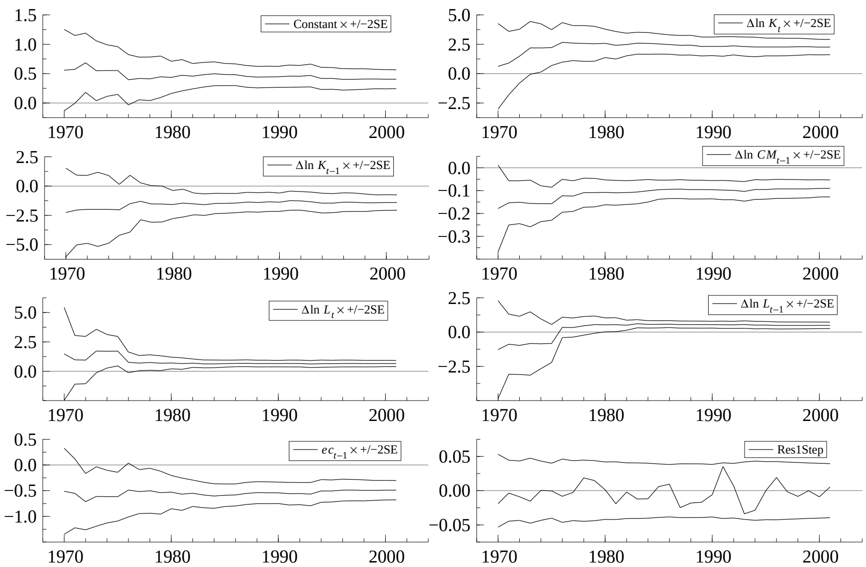


Figure 6: The recursive values of the parameter estimates along with the one-step residuals (Res1step) for the conditional model (8)

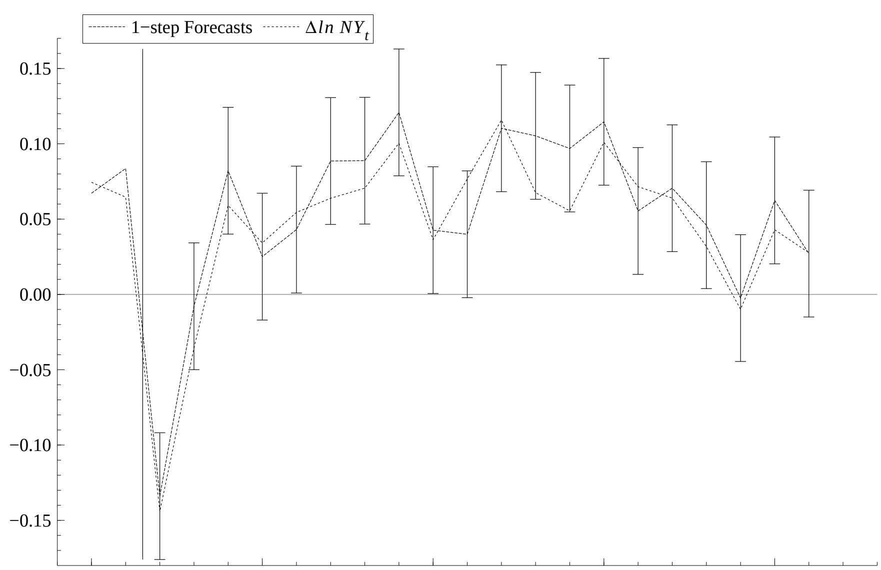


Figure 7: One-step ahead forecasts for the conditional model (8)

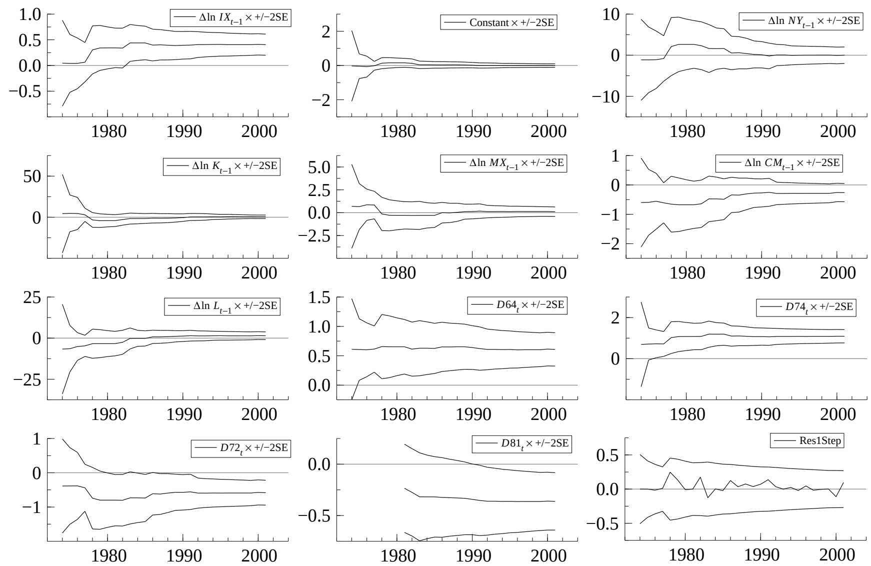


Figure 8: The recursive values of the parameter estimates along with the one-step residuals (Res1step) for the marginal model (9)

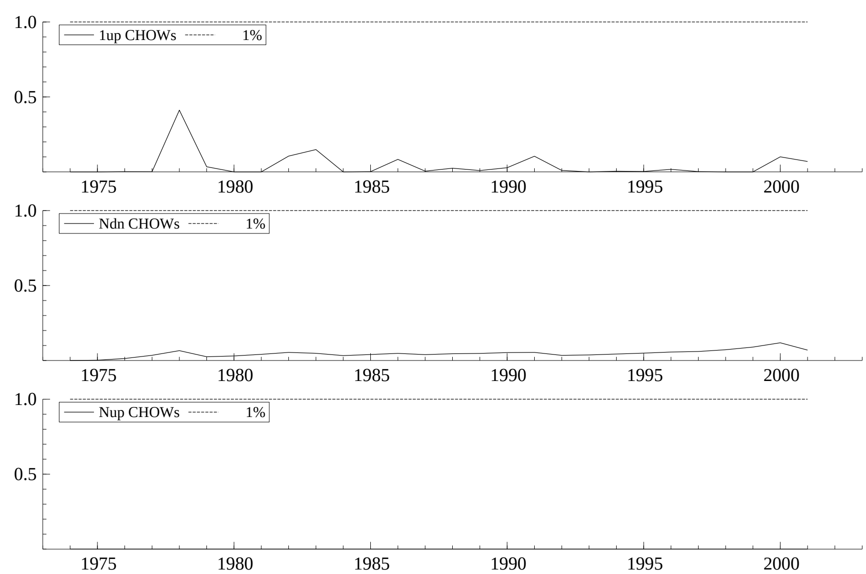


Figure 9: The recursive Chow test statistics for the marginal model (9) scaled by the corresponding 1% critical values